Trade Credit and Pricing: An Empirical Evaluation*

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Abstract

We empirically investigate the proposition that firms charge premia on cash prices in transactions involving trade credit. Using a comprehensive Swedish panel dataset on product-level transaction prices and firm-characteristics, we relate trade credit issuance to price setting. In a recession characterized by tightened credit conditions, we find that prices increase significantly more on products sold by firms issuing more trade credit, reflecting their larger exposures to increased funding costs and counterparty risks. Our results thus demonstrate the importance of trade credit for price setting and show that trade credit issuance induces a channel through which financial frictions affect prices.

Keywords: Trade credit; prices; inflation; liquidity; counterparty risk.

JEL: E31; E32; D22; G30; L11.

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I Introduction

Early theoretical work by Schwartz (1974) points to a neglected aspect of the firm's classical price-setting problem. Schwartz proposes the existence of a premium on the price of a good if trade credit is extended in the transaction. That is, if the transaction entails separation of delivery and payment in time, then a premium is added to the price set for a cash transaction. The premium is increasing in the maturity of the trade credit contract, the seller's funding costs, and the buyer's default risk. Since trade credit is an abundant feature of inter-firm trade and a significant part of firms' short-term financing, shifts in trade credit price premia—caused, for instance, by the sharp increases in financing costs and counterparty risks typically observed in periods of financial distress—could have a large impact on prices and make for an important channel through which financial frictions interact with prices.[1]

The purpose of this paper is to test the hypothesis that trade credit prices include a premium determined by the seller's funding costs and credit risk exposures. Our empirical evaluation is based on a dataset comprising product-plant level data on prices and quantities for all Swedish manufacturing firms above a certain size threshold, firm-level accounting data for the universe of Swedish corporations, and loan-level data covering all loans extended by the four major Swedish banks to Swedish corporations. These data allow us to relate firm-product inflation rates to firms' trade credit issuance, while carefully assessing the robustness of our results and validating the plausibility of our identifying assumptions. More specifically, our empirical design is geared to assess the influence of trade credit issuance on price-setting in the 2008–09 recession in Sweden. The recession—characterized by a severe credit crunch as well as a sharp downturn in the real economy—is of key importance for identification, since it led to widespread increases in funding costs and counterparty risks, while being caused primarily by external shocks hitting the

[1] Jacobson and von Schedvin (2015) show that the average amount of accounts receivable and payable, scaled by assets, are 16 and 11 percent for Swedish firms. Similar reliance on trade credit financing prevails across countries. For instance, Rajan and Zingales (1995) show that the corresponding numbers for a sample of US firms are 18 and 15 percent.

Swedish economy in the wake of the global financial crisis.

Our main finding can be summarized as follows. Firms that issued more trade credit, relative to firms that issued less, increased their prices significantly more in the 2008–09 recession. By comparing firms at the 10th percentile with firms at the 90th percentile of the pre-crisis trade credit issuance distribution, we find that the annual firm-product inflation rate in the 2008–09 recession is 2.9 percentage points higher for firms that issued more trade credit, which is substantial given that the mean of the annual price adjustment across firm-products over the full sample period amounts to 2.8 percent.

The assumptions underlying identification are validated in several ways. Firstly, there is no significant divergence in pre-treatment trends between low and high trade credit firms, which mitigates a concern that the documented effects would be present absent the 2008–09 recession. Secondly, we document that the positive relationship between trade credit issuance and price changes during the recession is larger for firms that faced higher increases in funding costs and counterparty risks, respectively, which indicates that our results are indeed associated with the mechanisms proposed in our conceptual framework. Thirdly, we show that the impact of trade credit issuance on prices remains present when we control for a broad set of factors that previous research has shown to be important determinants of firms' price-setting behaviour, which relaxes a concern that our results are outcomes of spurious correlations.

This paper demonstrates the relevance of a largely neglected aspect of the firm's price-setting problem: the pricing of trade credit. The previous literature has essentially concerned the setting of cash prices and has overlooked the possibility that firms charge premia to compensate for costs in trade credit issuance. Moreover, our results also contribute to the growing literature on the influence of financial frictions on price setting. Chevalier and Scharfstein (1996) and Gilchrist et al. (2017), for example, show that liquidity constraints may give rise to countercyclical markups, as firms raise prices to strengthen their liquidity positions in periods when liquidity is scarce. Our paper documents a complementary channel—the trade credit channel—through which financial frictions may affect prices.

The rest of the paper is organized as follows. The next section presents a conceptual framework that outlines the link between trade credit and pricing. Section 3 describes the 2008–09 recession in Sweden, details our data resources, and provides some descriptive statistics. The empirical framework is presented in Section 4 and the results are presented in Section 5. Section 6 concludes.

II Conceptual Framework

In standard formulations of the firm's price-setting problem, the optimal price for product p sold by firm i, $P_{i,p}$, is equal to the product of the firm's marginal cost for producing p, $MC_{i,p}$, and a mark-up, $\mu_{i,p}$, that depends on the firm's price-setting power in the product market:

$$P_{i,p} = \mu_{i,p} \cdot MC_{i,p}. \tag{1}$$

This characterization of the price-setting problem neglects one salient aspect, however, namely that inter-firm transactions ever so often involve trade credit. In trade credit transactions, sellers extend credit to buyers by allowing payment at a date later than that of delivery.[2]

Since lending is associated with costs—most importantly due to funding and to credit risk exposure—prices charged in trade credit transactions likely surpass prices charged in cash transactions. Schwartz (1974) highlights this trade credit feature of price-setting and posits that firms add a trade credit premium to the cash price, determined by contracted loan maturity and an implicit interest rate. Our conceptual framework—intended to support the subsequent empirical analyses—rests on the theoretical relationships discussed by Schwartz and we focus on their

[2]Trade credit contracts are usually formulated in net terms, which means that the buyer is required to pay within a specified period after delivery. The most common contracted maturity in Sweden is 30 days, but both shorter and longer periods are used. Giannetti, Burkart and Ellingsen (2011) and Klapper, Laeven and Rajan (2012) show that net-terms contracts are by far the most common in samples of American and European firms. An alternative to the net terms contract is the two-part contract, a well-known variety of which is the ”2/10 net 30,” which gives the buyer a discount of two percent for payments made within ten days of delivery, but no discount for payments made between eleven and thirty days after delivery.

implications for the link between firms' trade credit issuance and pricing decisions.

To formalize, let $P^C_{i,p}$ denote the cash price, corresponding to the price in Equation (1), and let $P^T_{i,p}$ denote the trade credit price. $P^T_{i,p}$ can then be expressed as a function of said cash price, the maturity, and an interest rate:

$$P^T_{i,p} = P^C_{i,p} \cdot e^{r_{i,p} \cdot \tau_{i,p}}, \tag{2}$$

where $r_{i,p}$ is the implicit annual interest rate charged by the seller for the trade credit loan and $\tau_{i,p}$ is the maturity of the trade credit contract, in number of net days divided by 365. The interest rate and maturity may well vary across transactions; the parameters $r_{i,p}$ and $\tau_{i,p}$ should therefore be interpreted as averages across all sales of product p by firm i.

From equations (1) and (2), we can derive the firm-product inflation rate:

$$\begin{aligned}\pi^T_{i,p,t} &= \ln P^T_{i,p,t} - \ln P^T_{i,p,t-1} \\ &= \Delta \ln \mu_{i,p,t} + \Delta \ln MC_{i,p,t} + \Delta \left(r_{i,p,t} \cdot \tau_{i,p,t} \right) \\ &= \Delta \ln \mu_{i,p,t} + \Delta \ln MC_{i,p,t} + \left(\tau_{i,p,t-1} \cdot \Delta r_{i,p,t} + r_{i,p,t-1} \cdot \Delta \tau_{i,p,t} \right).\end{aligned} \tag{3}$$

If we assume that maturities are approximately constant over time, implying that $\Delta\tau \approx 0$, the firm-product inflation rate in year t is determined by the change in the mark-up, the change in marginal cost, and the product of the average trade credit maturity and the change in the implicit interest rate; all of which are allowed to vary at the firm-product level. If firms, on the contrary, can adjust maturities in trade credit contracts in response to shifts in funding costs and counterparty risks, they may choose to shorten maturities when such shifts occur. This would attenuate the relationship between r and π^T, since some part of the direct effect of an increase in r on π^T would be offset by the decrease in τ. We return to the issue of how changes in trade credit contract maturities may affect our results in Section IV.B below.

Our hypothesis is that the trade credit interest rate r is determined primarily by two factors: (i) the seller's cost of funding the loan and (ii) the risk of default on the part of the customer. That is, the implicit interest rate underlying trade credit is increasing in sellers' funding costs, as well as in their credit risk exposure, all else equal. The funding cost, in turn, is determined by the shadow price of liquidity facing the firm—the opportunity cost of the marginal unit of liquidity—and is thus equal to the interest rate paid on short-term borrowing for firms that face no binding liquidity constraints, but higher than this for firms that do face binding liquidity constraints due to credit rationing.

It is unlikely that firms can set higher prices than their competitors for prolonged periods, since at some point customers will overcome switching frictions and turn to suppliers offering lower prices. Permanent differences in funding costs across firms should therefore not be reflected in corresponding price differences across firms. However, if a firm operates in a customer market, i.e., a market in which the customer base is sticky—for instance because of costly switching (Klemperer, 1987), costly search (Hall, 2008), or idiosyncratic preferences (Bronnenberg, Dubé and Gentzkow, 2012)—then its prices may differ from competitors' prices in the short-run. Phelps and Winter (1970) and Bils (1989) show that an important feature of price setting in customer markets is the trade-off between maximizing short-term revenue by increasing prices and building a future customer base by lowering prices. This suggests that it may be optimal for firms to pass on temporary variation in funding costs to buyers in trade credit transactions, in particular the sharp but temporary increases in funding costs that typically occur during financial crises.

The same reasoning is not necessarily true for the credit risk component, since any potential seller faces the same credit risk when extending trade credit to a given buyer and therefore will require the same actuarially fair compensation for bearing this risk—abstracting from differences in credit risk attributable to the terms in the trade credit contract. Hence, the credit risk component of the implicit interest rate may well reflect both temporary and permanent differences in counterparty risk faced by different sellers.

III Setting, Data, and Descriptive Statistics

A The 2008–09 recession in Sweden

We exploit the 2008–09 recession in Sweden to test the hypothesis that trade credit prices include a price premium determined by an implicit interest rate reflecting the seller's funding cost and credit risk exposure. The 2008–09 recession is well suited for this purpose, in featuring a sharp downturn in the real economy as well as severe distress in the banking sector—and since the origin of both lay in external shocks hitting the Swedish economy in the wake of the global financial crisis.[3]

The banking sector distress was largely due to two external shocks. Firstly, the collapse of international financial markets following the outbreak of the subprime crisis in the US. While Swedish banks had little direct exposure to mortgage-backed securities issued in the US, the Swedish banking sector is highly dependent on external wholesale funding and therefore sensitive to conditions on international financial markets. Secondly, the severe economic crisis in the Baltic countries caused large loan losses for two of Sweden's four major banks, which had expanded rapidly on the Baltic market prior to the crisis.[4] The Baltic crisis naturally affected the exposed banks more, but the unexposed banks were partly affected as well, since the problems stemming from the Baltic countries gave rise to concerns about the stability of the Swedish banking sector as a whole. These two shocks led to increased distress in the banking system, although observers' judgments differ somewhat as to the severity of the distress. According to the IMF's banking crisis database, for example, Sweden suffered a borderline systemic banking crisis beginning in 2008 (Laeven and Valencia, 2012), while Romer and Romer (2017), using a financial distress measure ranging from 0 to 15, classifies the level of distress in Sweden during 2008–09 as 5 on average, with a peak value of 7.

The banking sector distress quickly led to a deterioration in the credit conditions facing corporate borrowers: beginning in 2008 and continuing throughout 2009,

[3]See, e.g., Bryant, Henderson and Becker (2012) for a comprehensive discussion of the causes and consequences of the 2008–09 recession in Sweden.

[4]The Swedish bank market is dominated by four major banks, jointly accounting for around 85 percent of banking sector assets and 75 percent of corporate lending.

growth in bank lending to firms fell steadily (Finansinspektionen, 2012), and many firms reported on a worsening access to external finance (Sveriges Riksbank, 2009; Konjunkturinstitutet, 2009). Meanwhile, the real economy fell into a sharp recession, with a decline in real GDP of around six percent in 2009, partly due to the domestic banking sector distress and partly due to the breakdown in international trade, which hit the export-oriented Swedish economy badly and resulted in a doubling of aggregate bankruptcy rates. Thus, the events unfolding during the 2008–09 recession increased funding costs as well as credit risk exposures in the corporate sector; both of which yielded a rise in r, according to the hypothesis outlined in the previous section.

B Data and variable definitions

The empirical analysis in this paper is based on data from four sources, which we merge unambiguously by means of the unique identifier (*organisationsnummer*) attached to each Swedish firm. Firstly, we obtain data on prices and quantities from "Industrins varuproduktion," an annual survey conducted by Statistics Sweden comprising all manufacturing plants with at least 20 employees, as well as a sample of smaller plants. The data cover transaction prices and quantities of goods sold at the product-plant level; where products are classified using 8/9-digit CN codes.[5] Thus, for each product produced at a given plant, we observe the average transaction price (as opposed to the list price), as well as the quantity of goods sold in each year. We aggregate the price and quantity data to the firm-product level using the sales value for each product and plant as weights.

Secondly, we obtain firm-level accounting data from the database Serrano, which covers the universe of corporations in Sweden. Serrano is constructed based on data from several official sources, most importantly the Swedish Companies Registrations Office, to which all Swedish corporations are required to submit annual financial accounting statements in accordance with EU standards. Thirdly, we

[5] These data have previously been used by Carlsson and Skans (2012). To give an idea of the granularity of the product classification, we can, for example, consider the codes 84212100 and 84212200, which refer to 'machinery and apparatus for filtering or purifying water' and 'machinery and apparatus for filtering or purifying beverages (excl. water)', respectively.

use a loan-level database available at Sveriges Riksbank, which covers all loans and credit lines extended by the four major Swedish banks to Swedish corporations. Finally, we obtain data on firm-level default probabilities from UC AB, the leading credit bureau in Sweden.

Our primary outcome variable is the firm-product inflation rate, defined as the log change in average transaction prices for product p, charged by firm i, between $t-1$ and t:

$$\pi_{i,p,t} = \ln P_{i,p,t} - \ln P_{i,p,t-1}.$$

We obtain 49,134 firm-product inflation rate observations, corresponding to 3,928 firms and 3,917 unique products, over the sample period 2004–2011.[6] Panel A of Figure I shows the distribution of annual firm-product inflation rates. Around 16 percent of the observations are located in the ± 0.5 percent interval around zero, while around half are larger than 5 percent in absolute value.

The main explanatory variable concerns firms' trade credit maturities. For want of contract-level data and the exact maturity in each trade credit contract, we use the ratio of accounts receivable to sales divided by 365:

$$\hat{\tau}_i^{07} = \frac{Rec_i^{07}}{Sales_i^{07}} \cdot \frac{1}{365}.$$

$\hat{\tau}_i^{07}$ is thus a proxy for firm i's average trade credit maturity across all its products and customers in 2007.[7] We fix this variable to its last pre-crisis value to mitigate endogeneity concerns, but we confirm below that our results are robust to allowing τ to vary over time.

We include two sets of control variables. The first set of controls, $\mathbf{X_{i,p,t}}$, consists of two variables at the firm-product level: the log change in the quantity of sales of product p by firm i between years $t-1$ and t, $\Delta Q_{i,p,t}$; and the change in unit input costs for product p produced by firm i between years $t-1$ and t,

[6] The data contain several observations of very large price changes, which may well reflect unobserved changes in product quality. We remove such observations by truncating the inflation rate variable at the 5th and 95th percentiles.

[7] More precisely, $\hat{\tau}_i^{07}$ measures average time to payment, which may differ from contracted payment time due to either late or premature payments.

$\Delta UIC_{i,p,t}$. Unit input costs are defined as the sum of labor costs and intermediate input costs divided by physical output.[8] The second set of controls, $\mathbf{Z_{i,t-1}}$, comprises the following firm-level variables: cash and liquid assets, $Cash/Assets_{i,t-1}$; leverage, $Total\ debt/Assets_{i,t-1}$; asset tangibility, $Tangible\ assets/Assets_{i,t-1}$; cash flow, $EBITDA/Assets_{i,t-1}$; and firm size, $\ln Assets_{i,t-1}$. We winsorize the explanatory variables at the 1st and 99th percentiles to reduce the influence of outliers.

C Sample and descriptive statistics

Table I reports descriptive statistics for all variables used in the empirical analysis. The mean (median) firm-product inflation rate, reported in Panel A, is 2.8 (0.6) percent. The average value of $\hat{\tau}_i^{07}$, reported in Panel B, is 0.097, corresponding to a trade credit contract maturity of 35 days. Sixty percent of the firms have access to a non-exhausted credit line and the average size of the unused part is 4.3 percent of total assets. Panel C, finally, shows the values of the time-varying firm-characteristics. The average firm has a book value of assets of 283 million SEK and sales of 355 million SEK (roughly 44 and 55 million USD, respectively, at the exchange rate prevailing at the end of 2007). The sample thus consists primarily of medium and large firms.

In Panel B of Figure I, we show that our sample is representative of the broader economy in terms of price changes. More specifically, the figure shows that the average firm-product inflation rates in our sample tracks the changes in the aggregate producer price index for the goods-producing sector of the economy fairly closely over the entire sample period.

[8] We do not observe labor costs and intermediate input costs at the product level, so we must resort to the following approximation when calculating $\Delta UIC_{i,p,t}$. For each plant and year, we portion out plant-level labor costs and intermediate input costs across the products produced at the plant in proportion to the total plant production share of the product. We approximate the share of each product as the physical quantity produced during the year times the firm's average unit price for the product over the entire sample period. We define the shares in this way—instead of simply using the shares of each product in total sales at the plant—in order to avoid introducing spurious positive correlation between price changes and $\Delta UIC_{i,p,t}$. We then compute the plant-product level values of $\Delta UIC_{i,p,t}$. Finally, we aggregate $\Delta UIC_{i,p,t}$ to the firm-product level using total sales for each product-plant as weights.

IV Empirical Framework

A Empirical strategy

Our empirical strategy is to exploit an aggregate shock that generically increased funding costs and counterparty risks, and achieve identification using the cross-sectional variation in τ, which determines the sensitivity of firms' prices to such increases, cf. Equation (3). The identifying variation thus comes from the differences in average trade credit maturities across firms that prevailed at the time of the aggregate shock to r. Previous research has documented several factors that affect trade credit contract maturities, including financial factors (Garcia-Appendini and Montoriol-Garriga, 2013), product characteristics (Giannetti, Burkart and Ellingsen, 2011), market power (Klapper, Laeven and Rajan, 2012), and legislation (Barrot, 2016). An identifying assumption underlying our strategy—to be addressed in more detail below—is that this variation is uncorrelated with any unobserved factors that may have affected firms' price setting in the crisis.

The hypothesis that increases in funding costs and credit risk exposure cause firms to increase product prices can be tested using the following difference-in-differences specification:

$$\pi_{i,p,t} = \beta \cdot Crisis_t \cdot \hat{\tau}_i^{07} + \alpha_{i,p} + \alpha_t + \gamma \cdot \mathbf{X_{i,p,t}} + \delta \cdot \mathbf{Z_{i,t-1}} + \varepsilon_{i,p,t}. \tag{4}$$

where $\pi_{i,p,t}$ is the firm-product inflation rate; $Crisis_t$ is a dummy variable equal to one in the years 2008 and 2009 and zero otherwise; $\hat{\tau}_i^{07}$ is the average trade credit maturity for firm i in 2007; $\alpha_{i,p}$ and α_t are firm-product and year-fixed effects, respectively; and $\mathbf{X_{i,p,t}}$ and $\mathbf{Z_{i,t-1}}$ are the vectors of control variables defined in Section III.B. The firm-product fixed effects control for potential time-invariant differences in price setting between firms with low and high trade credit issuance, respectively, while the vector $\mathbf{X_{i,p,t}}$ controls for fluctuations in demand and production costs at the firm-product level. The vector $\mathbf{Z_{i,t-1}}$, finally, controls for additional time-varying firm-characteristics that may influence price setting. Standard errors are clustered at the firm-level in all regressions.

We estimate the baseline specification for the period 2004–2011, which comprises a four-year pre-crisis period (2004–2007), the crisis period itself (2008–2009), and a two-year post-crisis period (2010–2011). The coefficient of interest is β, measuring the extent to which trade credit issuance affected firm-product inflation rates in the crisis period.

Following the reasoning outlined in Roberts and Whited (2012), our empirical analysis rests on two identifying assumptions:

(A1) In the absence of the crisis, average price changes would have been the same across firms, irrespective of their degree of trade credit issuance.

(A2) There is no omitted variable correlated with trade credit issuance that affects prices during the crisis.

We assess the plausibility of these assumptions in the following ways. Firstly, we test for differences in pre-crisis trends in firm-product inflation rates between firms with low and high trade credit issuance, respectively. Secondly, we test for cross-sectional heterogeneity in the effects of trade credit issuance on firm-product inflation rates with respect to firms' pre-crisis liquidity positions, and to increases in counterparty risk during the crisis, respectively. If the relationship between trade credit issuance and prices during the crisis period can be attributed to increases in funding costs and counterparty risk, then this relationship should be stronger for liquidity-constrained firms, as well as for firms that faced larger increases in counterparty risk. Finally, we control for an additional set of potentially important confounding factors. For example, Chevalier and Scharfstein (1996) and Gilchrist et al. (2017) document that firms liquidity positions underlie countercyclical mark-ups. To control for such mechanisms, we estimate augmented versions of Equation (4) where we include interaction terms between the crisis variable and liquidity-related pre-crisis firm characteristics.

B The role of changes in trade credit contract maturities

A firm may, as noted above, respond to increases in funding costs and counterparty risks by raising the implicit interest rate on its trade credit lending, by reducing the maturity on this lending, or by some combination of the two. Any relationship between r and π^T will be attenuated if firms partly respond by lowering trade credit maturities, cf. Equation (1). We do not observe contracted trade credit maturities in our data, so the results in our empirical analysis concern the effects of r on π^T, net of any changes in τ. This means, that the more firms experiencing increases in funding costs or counterparty risks in the crisis respond by reducing contracted maturities—rather than increasing the trade credit interest rate—the smaller will the coefficient β in Equation (4) turn out. In the extreme, if all firms were to reduce maturities to zero—and effectively turn to cash transactions with trade credit premiums set to zero—our estimated effect should be negative. The relative importance of increases in interest rates and decreases in maturities is ultimately an empirical question, and we stress that our analysis captures the net effect on prices of these two potential responses.

While we cannot disentangle the effects of changes in interest rates and maturities in our empirical analysis, we believe that there are grounds for assuming that trade credit contract maturities are relatively sticky—and that our results thus primarily capture the direct effect of r on π^T. We provide some evidence supporting this assumption in Figure II, which shows average trade credit contract maturities, as measured by $\hat{\tau}_{i,t}$, within each tertile of the trade credit maturity distribution over the period 2004–2011. The figure in Panel A—in which we classify firms on a year-by-year basis, so that firms may switch between tertiles over time—shows that the the trade credit maturity distribution is very stable over time; the average maturity within each tertile hardly changes at all between 2004 and 2011, with the exception of a very small decrease during the crisis.

In the figure in Panel B, firms are classified according to their position in the trade credit distribution in 2007, so that the set of firms in each tertile is fixed over time. This figure shows that trade credit maturities changed very little for the firms in the bottom and middle tertiles, but that the firms in the top tertile reduced maturities somewhat during the crisis, with $\hat{\tau}_{i,t}$ falling from 0.136 to 0.114 (i.e., from around 50 to 42 days). Thus, while there may be some scope for firms to change the maturity of their trade credit lending in response to shocks, the evidence presented in Figure II is nevertheless consistent with the assumption that trade credit contract maturities are sticky.

V Results

A Main results

Figure III provides an illustration of our main finding. It shows average firm-product inflation rates over the period 2004–2011 for firms with average trade credit maturities above (solid line) and below (dashed line) the sample median in year $t-1$. Inflation rates for the two groups of firms track each other closely in the four years leading up to the crisis, but then differ substantially during the 2008–09 crisis period. Although average inflation rates fall in both groups of firms—which is what one would expect in a crisis period with deflationary pressure—inflation rates fall considerably less among firms with long trade credit maturities. In the post-crisis period, inflation rates resume similarity across the two groups of firms. Thus, the figure provides initial support for our hypothesis that increases in funding costs and counterparty risk lead firms to raise trade credit premia. We will next substantiate by means of a formal analysis using the model specified in Equation (4).

Table II reports the results for various estimations of the model specification in Equation (4). The baseline result is reported in Column (I). The coefficient on the interaction term, $Crisis_t \cdot \hat{\tau}_i^{07}$, is 0.205 and statistically significant, which implies that inflation rates during the crisis increased for firms with long trade credit maturities relative to firms with short maturities. The economic significance of this coefficient can be quantified by the difference in product-specific inflation rates between firms at the 90th and the 10th percentiles of the trade credit maturity distribution, which we find to be a substantial 2.9 percentage points.[9]

Next, we re-estimate the baseline specification using weights that adjust for differences in the shares of each firm's total sales accounted for by each of its products. More specifically, we estimate a weighted regression where the weight for each ob-

[9]When replacing the time-invariant explanatory variable, $\hat{\tau}_i^{07}$, with lagged, time-varying maturities, $\hat{\tau}_{i,t-1}$, we obtain a coefficient of 0.183 (3.1) for the baseline specification.

servation, $\omega_{i,p,t}$, is calculated as firm i's sales of product p firm divided by firm i's total sales. The results are reported in Column (II). The crisis dummy coefficient remains positive and statistically significant, but it is slightly smaller in magnitude than the coefficient in the baseline specification; the difference between firms with long and short trade credit maturities is now 2.0 percentage points. This suggests that firms are more prone to increase prices on non-core products when funding costs and counterparty risks increase.

In spite of winsorization, there is still a concern for undue influence from a small number of firms with exceptionally long trade credit maturities. We therefore estimate a version of the baseline specification in which the main explanatory variable is a dummy indicating whether a firm's trade credit maturity was above or below the sample median in the last pre-crisis year. The results, reported in Column (III), are consistent with the baseline results; the difference in inflation rates during the crisis between firms above and below the sample median of the trade credit maturity distribution is 1.3 percentage points. Similarly, one may be concerned that very large price adjustments drive the baseline result. To address the latter, we estimate a version of the baseline specification in which the dependent variable is replaced by a dummy that takes the value one for price increases, and zero otherwise. The estimated coefficient, reported in Column (IV), implies that firms with long trade credit maturities were 5.6 percentage points more likely to increase prices. These findings suggest that outliers in the dependent variable, or in the main explanatory variable, are not a concern for the baseline result.

The baseline specification includes firm-product fixed effects to control for time-invariant differences in inflation rates across products. Hypothetically, time-varying differences in inflation rates across products could be important. Supposing that inflation rates during the crisis were lower for certain products, for reasons unrelated to trade credit issuance, and that the same products are customarily sold with long trade credit maturities, then our baseline result could be spurious. To address this possibility, we estimate a specification in which we replace the firm-product fixed effects with product-year fixed effects to control for the part of the variation in the inflation rate common to all producers of a given product. The resulting coefficient, reported in Column (V), is positive and statistically significant, with a magnitude of around half of the baseline coefficient. This seems to suggest that our baseline result is partly associated with time-varying product-specific factors—rather than increases in the trade credit premia only. However, the inclusion of nine-digit product fixed effects interacted with year fixed effects eliminates much of the identifying variation, which may in itself decrease the magnitude of the coefficient. It is nevertheless reassuring, that even in this very strict specification the coefficient remains positive and statistically significant.

On a related note, we assess whether time-varying differences in inflation rates across industries could affect our results. We do this by estimating the baseline specification augmented with industry$\times$year fixed effects, where industries are defined using two-digit SNI/NACE codes. The results, reported in Column (VI), shows that the estimate of the coefficient β is virtually unchanged compared to the estimate in the baseline specification. Time-varying industry effects are therefore not a concern for the baseline results.

Next, we tackle the possibility that our result could reflect a shift in demand during the crisis—away from sellers with short trade credit maturities and toward sellers with long maturities—as a result of longer trade credit maturities becoming more valuable for liquidity-constrained buyers during crises. To evaluate the demand-shift explanation, we regress the change in the quantity of sold goods, $\Delta Q_{i,p,t}$, on the right-hand side of the baseline specification. The idea is that an upward shift in demand for goods sold by firms with long trade credit maturities should cause an increase in both prices and quantities. Column (VII) shows, however, that the coefficient in this specification is negative and insignificant, which speaks against the alternative explanation based on shifts in demand.

Finally, we evaluate the parallel trends assumption by testing for differences in inflation rates between firms with long and short trade credit maturities in each year of the sample period. We do this using the baseline specification supplemented with interactions of the key explanatory variable, $\hat{\tau}_i^{07}$, and year-fixed effects. The resulting β-coefficients are plotted in Figure IV, using 2004 as base year. The coefficients are insignificant and close to zero in all pre- and post-crisis years, but positive and statistically significant in the two crisis years, which provides support for the parallel pre-treatment trends assumption.

B Cross-sectional heterogeneity

The finding that extensive trade credit issuance cause larger increases in prices during the crisis should—according to the hypothesis outlined in the conceptual framework—be the result of some combination of increases in funding costs and in counterparty risks facing firms. To verify these mechanisms for our results, we conduct cross-sectional heterogeneity analyses in which we estimate the baseline specification on sub-samples of firms. The sample-splits are defined by empirical approximations of changes in funding costs and counterparty risk in the crisis. Our conjecture is that the association between trade credit issuance and price changes is stronger for firms that experienced larger increases in funding costs and counterparty risk, respectively.

We approximate for changes in funding costs using two measures of firms' pre-crisis liquidity positions: cash and liquid assets, $Cash/Assets^{07}_{i,t-1}$; and the size of unused credit lines, $Unused\ LC/Assets^{07}_{i,t-1}$.[10] Firms with weaker pre-crisis liquidity positions were presumably more vulnerable to the deterioration in access to external finance during the crisis, and can consequently be expected to have experienced larger increases in funding costs. Changes in counterparty risk are approximated using industry-level measures of changes in the average default probabilities in customer industries during the crisis, $\Delta CP\ Risk^{07-09}_{j}$, where j denotes two-digit SNI/NACE industries.[11] For each variable, we construct two sub-samples: one with the firms in the bottom three deciles and one with the firms in the top three deciles.[12]

[10] We observe lending from the four major banks, which may lead us to underestimate $Unused\ LC/Assets^{07}_{i,t-1}$ for firms obtaining credit from minor Swedish banks. This will lead us to underestimate the difference between the two sub-samples, if anything.

[11] The counterparty-risk measure is constructed as follows. First, we calculate the sales-weighted average default probability (PD) for each two-digit SNI/NACE industry and year. We then compute the change in each industry's weighted average PD between 2007 and 2009. Finally, we use the 2008-vintage of Statistics Sweden's input-output tables to calculate the industry-level measure of changes in average customer PDs facing firms in each industry.

[12] The results are very similar if we instead define the sub-samples by splitting the sample at the

The results of the cross-sectional heterogeneity analyses are reported in Table III. Columns (I) and (II) cover the results for the sample splits based on the $Cash/Assets^{07}_{i,t-1}$-distribution. The coefficient is large and statistically significant for firms with low cash holdings, but relatively small and statistically insignificant for firms with high cash holdings; the difference is significant at the five-percent level. A similar pattern emerges in Columns (III) and (IV), where we report the results for sub-samples of firms with credit lines in the bottom and top of the $Unused\ LC/Assets^{07}_{i,t-1}$-distribution: the coefficient is large and significant in the former group, but smaller and insignificant in the latter; the difference is not statistically significant at the five-percent level in this case, however. The results reported in Columns (I)-(IV) thus support the notion that increases in funding costs account for some part of the positive relationship between trade credit issuance and price changes during the crisis.

The results for the sample splits based on changes in counterparty risk are reported in Columns (V) and (VI). The estimated coefficient is large and statistically significant in the sub-sample of firms that faced larger increases in counterparty risk during the crisis, but small and statistically insignificant in the group of firms for which the risk increase was smaller. The difference between the two coefficients is, moreover, statistically significant, which suggests that increases in counterparty risk contribute to the positive relationship between trade credit issuance and price changes during the crisis. The results concerning counterparty risk should be interpreted with some caution, however. There are two main reasons for this. Firstly, $\Delta CP\ Risk^{07-09}_{j}$ is measured at the two-digit industry level and is by construction a crude approximation for the change in counterparty risk facing an individual firm. Secondly, $\Delta CP\ Risk^{07-09}_{j}$ is an ex post measure that in principle could be subject to reverse causality; this would be the case if price increases by suppliers during the crisis—undertaken for reasons other than increased buyer credit risk—caused increased default risk for their buyers, rather than the other way around. These caveats notwithstanding, we believe that our results provide support for the hypoth-

median of each variable. The results from this alternative specification are available from the authors upon request.

esis that part of our baseline results is accounted for by increased counterparty risk.

In sum, the results reported in Table III suggest that the positive relationship between trade credit issuance and price changes during the crisis is related to increases in funding costs and in counterparty risks.

C Liquidity constraints and trade credit issuance

In a final exercise, we demonstrate the relevance of the mechanism in Chevalier and Scharfstein (1996) and Gilchrist et al. (2017) in our empirical setting, and show that it operates side-by-side with the mechanism proposed in this paper. We do this by estimating models intended to capture the direct effect of liquidity constraints on firms' price setting in the crisis. More specifically, the models are based on the baseline specification augmented with an interaction term between a liquidity-constraint proxy and the crisis variable:

$$\begin{aligned}\pi_{i,p,t} = \beta \cdot Crisis_t \cdot \hat{\tau}_i^{07} + \phi \cdot Crisis_t \cdot LC_i^{07} + \\ \alpha_{i,p} + \alpha_t + \gamma \cdot \mathbf{X_{i,p,t}} + \delta \cdot \mathbf{Z_{i,t-1}} + \varepsilon_{i,p,t}.\end{aligned} \tag{5}$$

We estimate Equation (5) using two proxies for liquidity constraints: pre-crisis leverage, $Total\ debt/Assets_i^{07}$, and pre-crisis cash holdings, $Cash/Assets_i^{07}$.

We also estimate the following variation on Equation (5), in which we estimate separate effects for each year of the crisis:

$$\begin{aligned}\pi_{i,p,t} = \beta^{08} \cdot Crisis_t^{08} \cdot \hat{\tau}_i^{07} + \beta^{09} \cdot Crisis_t^{09} \cdot \hat{\tau}_i^{07} + \\ \phi^{08} \cdot Crisis_t^{08} \cdot LC_i^{07} + \phi^{09} \cdot Crisis_t^{09} \cdot LC_i^{07} + \\ \alpha_{i,p} + \alpha_t + \gamma \cdot \mathbf{X_{i,p,t}} + \delta \cdot \mathbf{Z_{i,t-1}} + \varepsilon_{i,p,t}.\end{aligned} \tag{6}$$

The results of the estimations are reported in Table IV. Columns (I) and (II) show the results for the estimations where the liquidity-constraint proxy is pre-crisis leverage. The estimated coefficient on the interaction between leverage and the crisis variable, reported in Column (I), is positive and statistically significant, which implies that firm-product inflation rates during the crisis were higher for firms that entered the crisis highly leveraged. The coefficients reported in Column (II) show that this effect is present in both years of the crisis.

Columns (III) and (IV) show the results for the estimations where the liquidity-constraint proxy is pre-crisis cash holdings. The estimated coefficient on the interaction between cash holdings and the crisis variable, reported in Column (III), is negative but statistically insignificant. The yearly coefficients reported in Column (IV) show, however, that the effect of cash holdings on prices is significant in the first year of the crisis, which implies that firm-product inflation rates during the first year of the crisis were higher for firms that entered the crisis with lower cash holdings.

The results reported in Table IV thus show that the mechanism documented by Chevalier and Scharfstein (1996) and Gilchrist et al. (2017) is also in operation in our empirical setting. Nevertheless, the effect of interest in this paper—that of trade credit issuance on prices during the crisis—remains positive and statistically significant in all estimations. This shows that the effects documented in this paper are complementary to those documented in Chevalier and Scharfstein (1996) and Gilchrist et al. (2017).

In an unreported exercise, we also estimate a variation on the model in Equation (5) in which we interact *all* firm-level control variables in the baseline specification—rather than only a liquidity-constraint proxy—with the crisis dummy. We do this to test whether our result could be driven by some other firm characteristic that only affects prices in crisis periods and that is correlated with trade credit issuance. Our estimate of the coefficient β using this specification is 0.155 (2.9), which is slightly smaller but still close to the estimate from the baseline specification.

VI Conclusions

Theoretical research has proposed the existence of a trade credit price premium, governed by an implicit interest rate determined by the selling firm's funding costs and the buying firm's default risk. This implies that increases in funding costs and counterparty risks should generate larger impacts on inflation rates for products sold by firms that extend more trade credit. By means of a difference-in-differences approach applied to Swedish manufacturing firm data, we relate adjustments in firm-product inflation rates in the 2008–09 recession to pre-crisis trade credit issuance, towards an appraisal of the trade credit price premium hypothesis. We confirm that firms issuing more trade credit exhibited substantially higher adjustments in firm-product inflation rates during the crisis. The documented effects are stronger for liquidity-constrained firms—which supports the notion that they reflect increases in firms' valuation of liquidity caused by contractions in the availability of external financing—as well as for firms whose customers undergo downward shifts in creditworthiness. Hence, we find empirical support for the hypothesized determinants of the implicit interest rate in the trade credit price premium: funding costs and counterparty risks.

Our results contribute to the growing literature on the influence of financial market imperfections on firms' price setting. Notable work by Chevalier and Scharfstein (1996) and Gilchrist et al. (2017) show that liquidity constraints lead to countercyclical price mark-ups. Our paper highlights that trade credit issuance induces an additional channel—partly in parallel, but also over and above the previously documented one—that can explain countercyclical movements in mark-ups. But, more broadly, our paper demonstrates the relevance of an aspect of the firm's pricing problem neglected in the previous literature—which has essentially concerned itself with the setting of cash prices, and overlooked the firm's need to compensate for issued trade credit.

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Table I: Descriptive statistics

This table reports descriptive statistics for all variables used in the empirical analyses, as well as for some additional firm characteristics. Definitions of the variables are provided in the text.

	Mean	Median	Std. dev	Pct. 10	Pct. 25	Pct. 75	Pct. 90	No. obs
Panel A. Price and quantity variables (2004–2011)								
Firm-product inflation ($\pi_{i,p,t}$)	0.028	0.006	0.161	–0.149	–0.032	0.092	0.224	49,134
Change in quantity sold ($\Delta Q_{i,p,t}$)	–0.008	0.005	0.456	–0.433	–0.150	0.156	0.395	49,134
Change in unit input costs ($\Delta UIC_{i,p,t}$)	0.037	0.022	0.350	–0.254	–0.075	0.137	0.322	49,134
Panel B. Key explanatory and sample split variables (2007)								
Trade credit maturity ($\hat{\tau}_i^{07}$)	0.097	0.094	0.058	0.022	0.061	0.125	0.162	3,928
$Cash/Assets_i^{07}$	0.086	0.024	0.130	0.000	0.002	0.116	0.272	3,928
LC_i^{07} (0/1)	0.604	1.000	0.489	0.000	0.000	1.000	1.000	3,928
$Unused\ LC/Assets_i^{07}$	0.043	0.003	0.071	0.000	0.000	0.063	0.136	3,928
Panel C. Other firm characteristics (2004–2011)								
Trade credit maturity ($\hat{\tau}_{i,t-1}$)	0.090	0.089	0.049	0.025	0.059	0.117	0.149	18,885
$Cash/Assets_{i,t-1}$	0.081	0.022	0.125	0.000	0.002	0.110	0.258	18,885
$Total\ debt/Assets_{i,t-1}$	0.127	0.030	0.165	0.000	0.000	0.233	0.388	18,885
$Tangible\ assets/Assets_{i,t-1}$	0.267	0.245	0.184	0.040	0.112	0.392	0.527	18,885
$Cash\ flow/Assets_{i,t-1}$	0.075	0.074	0.140	–0.071	0.013	0.146	0.234	18,885
$Assets_{i,t-1}$ (in SEK 1,000)	282,522	57,417	796,530	14,022	25,702	159,289	539,728	18,885
$Sales_{i,t-1}$ (in SEK 1,000)	354,555	98,216	817,508	26,419	44,393	249,974	758,715	18,885

Table II: Baseline results

This table reports results for estimations of various specifications based on Equation (4). The dependent variable is the firm-product inflation rate, $\pi_{i,p,t}$, in all specifications except those in Column (IV), where the dependent variable is a dummy equal to one for price increases and zero otherwise, and in Column (VII), in which the dependent variable is the change in the quantity of goods sold, $\Delta Q_{i,p,t}$. The regression in Column (II) is estimated using WLS, where the weight for each observation, $\omega_{i,p,t}$, is calculated as firm i's sales of product p divided by firm i's total sales in year t. The product-fixed effects are based on 8/9-digit CN codes. The estimation period is 2004–2011 in all columns. t-statistics calculated using robust standard errors clustered at the firm-level are reported in parentheses. ***, **, and * denotes statistical significance at the 1, 5, and 10 percent levels, respectively.

	(I)	(II)	(III)	(IV)	(V)	(VI)	(VII)
Dependent variable:	$\pi_{i,p,t}$	$\pi_{i,p,t}$	$\pi_{i,p,t}$	$\pi^{+}_{i,p,t}$	$\pi_{i,p,t}$	$\pi_{i,p,t}$	$\Delta Q_{i,p,t}$
$Crisis_t \cdot \hat{\tau}_i^{07}$	0.205***	0.144**		0.400**	0.109**	0.208***	–0.124
	(3.6)	(2.5)		(2.5)	(2.1)	(4.2)	(–0.8)
$Crisis_t \cdot \hat{\tau}_i^{07,High}$			0.013**				
			(2.1)				
Economic significance	0.029	0.020	-	0.056	0.015	0.029	-0.017
Year FE	Yes	Yes	Yes	Yes	No	Yes	Yes
Firm × Product FE	Yes	Yes	Yes	Yes	No	Yes	Yes
Product × Year FE	No	No	No	No	Yes	No	No
Product and firm controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry × Year FE	No	No	No	No	No	Yes	No
Weights	No	Yes	No	No	No	No	No
R^2	0.327	0.398	0.327	0.302	0.556	0.352	0.340
Number of firms	3,928	3,928	3,928	3,928	3,928	3,928	3,928
Number of observations	49,134	49,134	49,134	49,134	49,134	49,134	49,134

Table III: Cross-sectional heterogeneity

This table reports results for estimations of the baseline specification in Equation (4) on various sub-samples of firms. Columns (I) and (II) report results for sample splits based on cash holdings; Columns (III) and (IV) for sample splits based on unused credit lines; and Columns (V) and (VI) for sample splits based on changes in counterparty risks. The cutoffs used to construct the sub-samples are defined at the firm level; hence, the number of firms in each sub-sample is approximately the same, while the number of observations differ somewhat. For the sample splits based on unused credit lines and changes in counterparty risk, the number of firms and observations also differ due to bunching of observations; in the former case because more than 30 percent of firms have no unused credit line, and in the latter because the cutoffs are defined using an industry-level measure. Reported p-values correspond to one-tailed tests, where the null hypothesis is that the β-parameters are equal in each pair, and the alternative hypothesis that the coefficients are larger in the groups of firms with low cash holdings, low credit lines, and high increases in counterparty risk, respectively. The estimation period is 2004–2011 in all columns. t-statistics calculated using robust standard errors clustered at the firm-level are reported in parentheses. ***, **, and * denotes statistical significance at the 1, 5, and 10 percent levels, respectively.

	Dependent variable: $\pi_{i,p,t}$					
	$Cash/Assets_i^{07}$		$Unused\ LC/Assets_i^{07}$		$\Delta CP\ Risk_j^{07-09}$	
	(I)	(II)	(III)	(IV)	(V)	(VI)
	Low	High	Low	High	Low	High
$Crisis_t \cdot \hat{\tau}_i^{07}$	0.365***	0.082	0.223***	0.081	0.058	0.452***
	(3.8)	(0.8)	(3.1)	(1.0)	(0.5)	(3.9)
p-value	0.020		0.094		0.007	
R^2	0.147	0.185	0.169	0.163	0.177	0.161
Number of firms	1,179	1,179	1,557	1,179	636	1,223
Number of observations	16,588	11,836	17,484	14,046	6,318	14,056

Table IV: Liquidity constraints and trade credit issuance

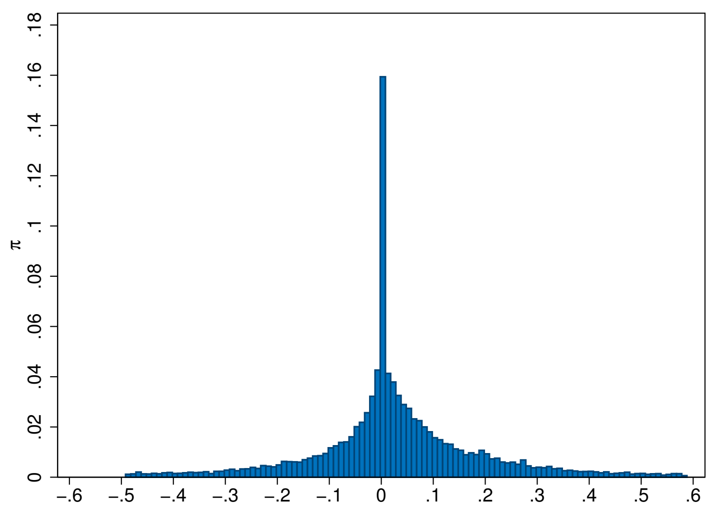
This table reports results for estimations of the models specified in Equations (5) and (6). The dependent variable is the firm-product inflation rate, $\pi_{i,p,t}$, in all specifications. The proxy for liquidity constraints is leverage in Columns (I) and (II) and cash-holdings in Columns (III) and (IV). The estimation period is 2004–2011 in all columns. t-statistics calculated using robust standard errors clustered at the firm-level are reported in parentheses. ***, **, and * denotes statistical significance at the 1, 5, and 10 percent levels, respectively.

	(I)	(II)	(III)	(IV)
	$LC_i^{07} = Total\ debt/Assets_i^{07}$		$LC_i^{07} = Cash/Assets_i^{07}$	
$Crisis_t \cdot \hat{\tau}_i^{07}$	0.190***		0.205***	
	(3.4)		(3.6)	
$Crisis_t \cdot LC_i^{07}$	0.043**		–0.014	
	(2.4)		(–0.6)	
$Crisis_t^{08} \cdot \hat{\tau}_i^{07}$		0.196***		0.206***
		(2.9)		(3.0)
$Crisis_t^{09} \cdot \hat{\tau}_i^{07}$		0.183**		0.203***
		(2.3)		(2.6)
$Crisis_t^{08} \cdot LC_i^{07}$		0.034*		–0.059**
		(1.7)		(–2.2)
$Crisis_t^{09} \cdot LC_i^{07}$		0.050**		0.035
		(2.0)		(1.1)
Year FE	Yes	Yes	Yes	Yes
Firm x Product FE	Yes	Yes	Yes	Yes
Product x Year FE	No	No	No	No
Product and firm controls	Yes	Yes	Yes	Yes
Industry x Year FE	No	No	No	No
Weights	No	No	No	No
R^2	0.327	0.327	0.327	0.327
Number of firms	3,928	3,928	3,928	3,928
Number of observations	49,134	49,134	49,134	49,134

Figure I: Firm-product inflation rates

Panel A of this figure shows the distribution of firm-product inflation rates in our sample over the entire sample period, 2004–2011. Panel B shows the average firm-product inflation rates in our sample, as well as the annual changes in the aggregate producer price index for the entire goods-producing economy (SNI/NACE sections A-E), for each year between 2004 and 2011. We calculate the latter as the log change in the annual average of the monthly values of the producer price index. *Source:* Statistics Sweden and authors' calculations.

A. Distribution of firm-product inflation rates in sample



B. Price changes in sample and in the aggregate economy

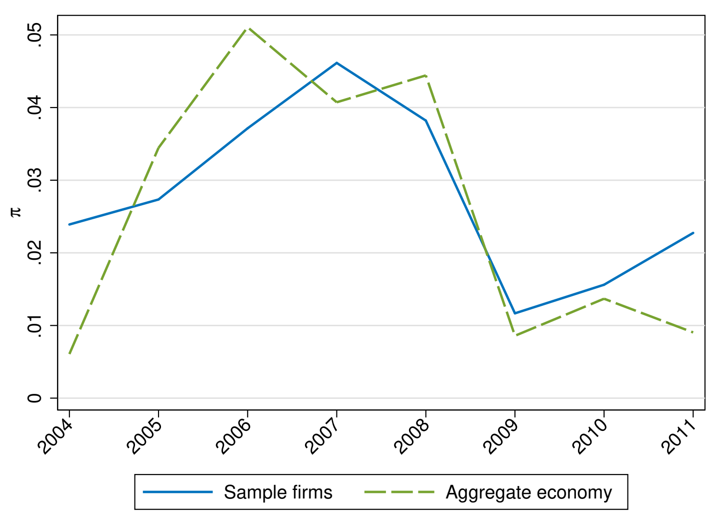
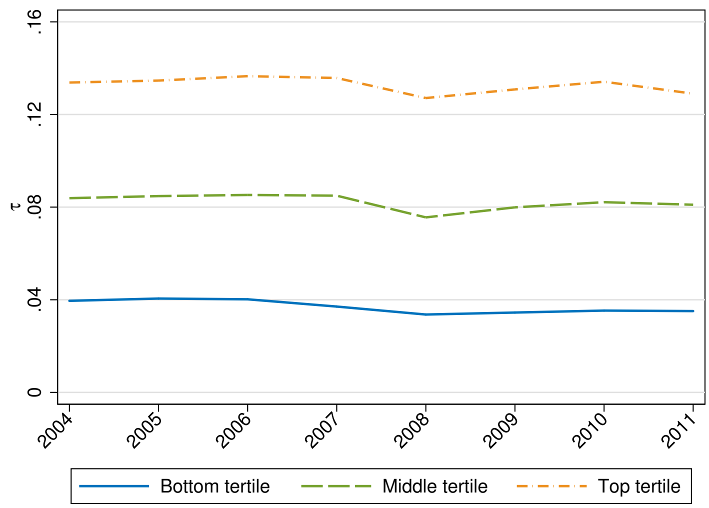


Figure II: Average trade credit maturity by tertile

This figure shows average trade credit maturities, as measured by $\hat{\tau}_{i,t}$, within each tertile of the trade credit maturity distribution. The figure in Panel A shows firms classified on a year-by-year basis, while in the figure in Panel B firms are classified according to their position in the trade credit maturity distribution in 2007.

A. Year-by-year classification of firms into tertiles



B. Fixed classification of firms into tertiles

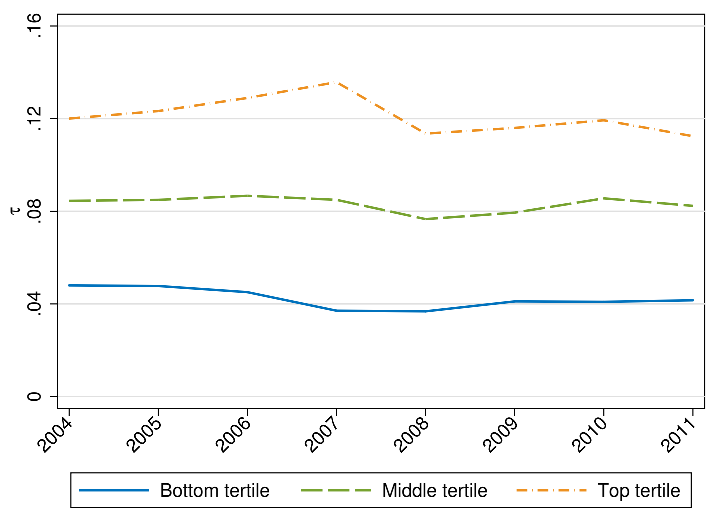


Figure III: Average firm-product inflation rates over time

This figure shows average firm-product inflation rates in each year of the sample period for firms above (solid line) and below (dashed line) the median of the trade credit issuance distribution in year $t-1$, $\hat{\tau}_{i,t-1}$.

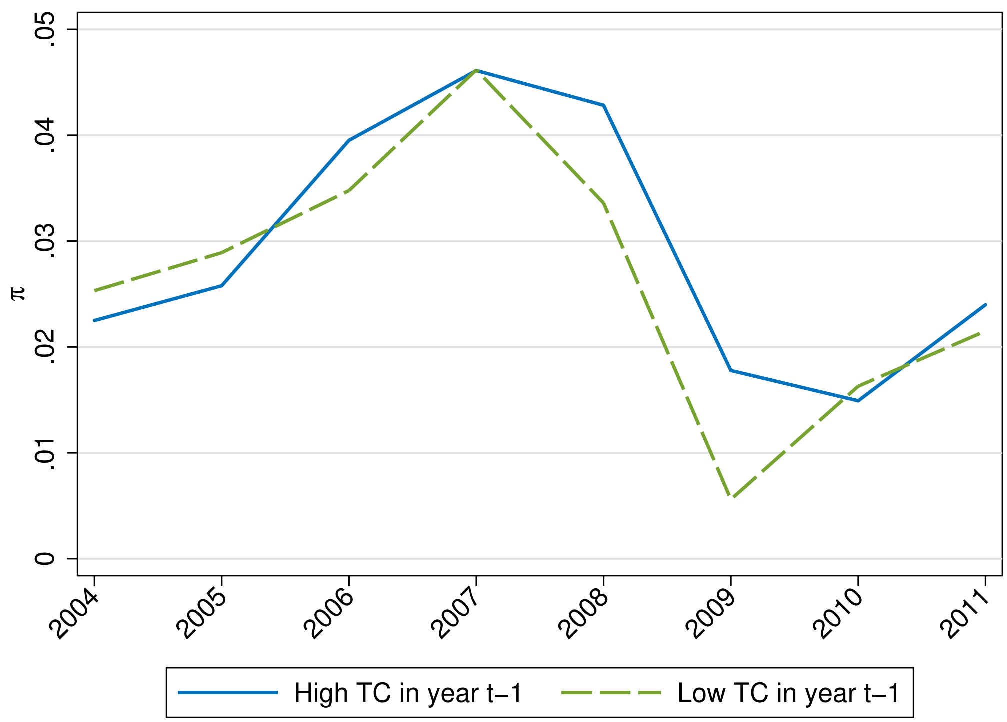


Figure IV: Pre-treatment trends

This figure shows the β_t-coefficients from an estimation of the baseline specification supplemented with interactions of the key explanatory variable, $\hat{\tau}_i^{07}$, and year fixed effects. The estimation is carried out using the entire sample period 2004–2011, with 2004 serving as base year. The estimating equation is thus: $\pi_{i,p,t} = \sum_{t=2005}^{2011} \beta_t \cdot Year_t \cdot \hat{\tau}_i^{07} + \alpha_{i,p} + \alpha_t + \gamma \cdot \mathbf{X_{i,p,t}} + \delta \cdot \mathbf{Z_{i,t-1}} + \varepsilon_{i,p,t}$. The vertical bars represent 95 percent confidence intervals.